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Job Search with Nonparticipation

Paul Frijters and Bas van der Klaauw

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*Australian National University, Research School of Social Sciences.
†Free University Amsterdam, Scholar and CEPR.
Address: Department of Economics, Free University Amsterdam, De Boelelaan 1105, NL–1081 HV Amsterdam, The Netherlands.

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1 Introduction

European labor markets are characterized by a low inflow into unemployment, long benefits entitlement and high proportions of long-term unemployed workers (e.g. Bean, 1994; Ljungqvist and Sargent, 1998; and Machin and Manning, 1999). For many possible reasons, individuals might quit searching for work while being unemployed and leave the labor force. Atkinson and Micklewright (1991) for instance note that in Germany around 30% of the unemployed workers leave to nonparticipation. Although this proportion is substantial, in the empirical microeconomic literature not much attention has been given to nonparticipation.

In the empirical literature that does consider nonparticipation, the usual approach is to consider transitions from unemployment into nonparticipation as stochastic occurrences. Burdett, Kiefer, Mortensen and Neumann (1984) analyze a model where individuals in a particular labor market state randomly receive offers from the other labor market states. Individuals accept an offer if it improves their expected present value of future utility. Such a stationary model is estimated by Mortensen and Neumann (1984). Similarly, Weiner (1984) estimates a reduced-form model allowing for duration dependence where nonparticipation is a competing risk. A related approach is chosen by Van den Berg (1990b) who estimates a (stationary) job search model, where each unemployed worker is exposed to the same risk of becoming nonparticipant.

These studies have in common that they assume that transitions into nonparticipation are only possible at stochastically determined moments. If a worker becomes disabled, considering nonparticipation as a risk seems reasonable. In case a worker is offered the possibility of early retirement, there arguably is both a stochastic and a choice element in the transition to nonparticipation. However, these reasons for nonparticipation mainly apply to employed workers. For unemployed workers, early retirement is impossible and the risk of disability is much smaller than for employed workers. Many unemployed workers who stop searching for jobs start a family or pursue different interests in life. This would seem to be a choice. Therefore, it is more appropriate to assume that at any moment during a spell of unemployment it is possible for an individual to quit searching for work and become nonparticipant voluntarily.

In this paper we allow unemployed workers to permanently have the possibility to quit searching for work and leave the labor force. In particular, we extend the job search framework of Van den Berg (1990a) by assuming that unemployed workers have this option. In a stationary environment, i.e. when both the expected returns to job search and the utility of being nonparticipant are constant, unemployed individuals either enter nonparticipation immediately after becoming unemployed or never become nonparticipant. Since we observe individuals moving into nonparticipation after some time spent in unemployment, the environment faced by the unemployed worker must be non-stationary. For the unemployed workers who enter nonparticipation, the expected present value of being unemployed must be declining in the unemployment duration. Additionally, because reservation wages include the value of future options, the permanent option of entering nonparticipation changes the reservation wage path and alters the distribution of accepted wages and re-employment rates. The job search model considered in this paper allows two structural elements to be non-stationary, i.e. the job offer arrival rate and the wage offer distribution. The structural approach taken in this paper allows us to distinguish between non-stationarity in these structural elements, and the observed transitions to nonparticipation improve their identification.

When studying the transition from unemployment to employment, failing to take exits to nonparticipation into account may cause several biases. First, unemployed workers with low expected returns to job search quit job searching earlier than unemployed workers with high expected returns of job search. Exits to nonparticipation are thus not independent of the re-employment probabilities and can therefore not be treated as exogenous right-censoring of the duration of unemployment. Second, since individuals with high expected returns to job search continue searching for a longer duration (e.g. Van Soest, Fontein and Euwals, 1996), these individuals are more likely to exit to employment.

In the theoretical literature many sources of non-stationarity in the job search process are discussed, and these sources are the foundations for many macroeconomic models of hysteresis in European unemployment levels. According to Bean (1994), the high level of European unemployment is mainly associated with low re-employment probabilities, rather than a high incidence of unemployment. This high occurrence of long-term unemployment stresses the importance of non-stationarity in the job search process. Blanchard and Diamond (1994) introduce ranking as an explanation for stigmatization of long-term unemployed workers. If employers use the unemployment duration of job applicants as a signal for their ability, long-term unemployed workers get stigmatized and receive less job offers. Along with the job offer arrival rate also the worker's reservation wage decreases, which affects post-unemployment wages. Calmfors and Lang (1995) provide discouragement of unemployed workers as a possible explanation for negative duration dependence. Discouragement has similar implications as stigmatization. Calmfors and Lang (1995) also mention the possibility that unemployed workers eventually get so discouraged by failing to find a job that they quit searching and choose to become nonparticipant permanently. Ljungqvist and Sargent (1998) argue that the persistently high European unemployment level in the last decades is due to a combination of institutional factors and (heterogeneous) loss of skills at the start and during unemployment. The generous unemployment benefits systems in most European countries causes the unemployed workers with few skills to be very selective in accepting job offers, which increases their unemployment duration. If unemployed workers are exposed to loss of skills, the generous unemployment benefits payments in Europe aggravate, especially in volatile economic periods, the percentage of long-term unemployed workers. In our model, loss of skills are mainly reflected by decreasing wages associated with job offers. See Machin and Manning (1999) for a survey of explanations of duration dependence and Blanchard and Wolfers (2000) for a survey of the consequences of negative duration dependence in the job search process for European unemployment rates.

Each source of duration dependence in the job search process has different policy implications. For example, training and schooling programs can be useful in case of loss of skills, extensive monitoring of search behavior when discouragement is the main source of duration dependence, and employer subsidies for hiring long-term unemployed might be a useful policy in the presence of stigmatization.[1] Therefore it is important to have an indication about the underlying source of duration dependence. Reduced-form empirical analyses cannot distinguish between different reasons for non-stationarity in the job search process. Machin and Manning (1999) mention that nearly all empirical work is based on the Mixed Proportional Hazard (MPH) framework.[2] Both Machin and Manning (1999) and Van den Berg (2001) argue that the MPH framework is convenient, but that it has no structural base. Only in very special cases, such as when unemployed workers are myopic or when job offers are never rejected, does the MPH framework have a structural interpretation. These cases are so restrictive that they are not very interesting to investigate. We will compare reduced-form estimation results with the estimation results of the structural model.

In the empirical analyses we use data from the German Socio-Economic Panel (GSOEP), which is a household survey. From this database we extracted a flow sample of individuals who entered unemployment between January 1989 and

[1]It is not clear if employer bonuses may overcome stigmatization of long-term unemployed workers. These bonuses might be a reasons for stigmatization itself.

[2]Recently, some studies focussed on the importance of 'scarring' due to unemployment on wage changes. Gregory and Jukes (2001) for instance estimate that in the UK a year of unemployment causes a 10% reduction in wages compared to a previous job. Their study does not address the underlying causes of this wage reduction, which makes interpretation difficult.

December 1995. Since we are mainly interested in reasons for leaving the labor force other than retiring, we only include unemployed workers younger than 45 years in our data set. Other than the length of the unemployment spell, the destination state and the post-unemployment wage, the database also includes the reservation wage. For some individuals we observe multiple unemployment spells. This allows us to distinguish true duration dependence in the job offer arrival rate and the wage offer distribution from unobserved and observed individual heterogeneity.

The procedure to estimate the model is non-standard. If the destination state after a spell of unemployment is nonparticipation, we know that at the moment of exiting unemployment the instantaneous utility of nonparticipation equals the reservation wage. However, if the destination state is employment (or the individual is still unemployed at the end of the observation period), we do not have any measure for the utility of nonparticipation. To compute the likelihood function we therefore have to integrate over all possible values of the utility of nonparticipation. Since this integral does not have a closed-form solution, we use simulated maximum likelihood to obtain parameter estimates (see Börsch-Supan and Hajivassiliou, 1994; Gourieroux and Montfort, 1996; and Stern, 1997).

The outline of the paper is as follows. Section 2 describes the non-stationary job search model which allows for endogenous nonparticipation. We derive the model and show some of its features. Furthermore, we briefly discuss the identification of the model and the parameterization. In Section 3 we briefly discuss some institutional aspects of German unemployment and give an overview of the data used in the empirical analyses. Section 4 presents the results of reduced-form analyses. In particular, we estimate a competing risks model for the transitions from unemployment to employment and to nonparticipation and we estimate a wage equation. The estimation results of the structural model are presented in Section 5, along with a discussion of the estimation methods. Section 6 concludes.

2 The non-stationary job search model

2.1 Outline of the model

In this section we present the structural framework used to model the transition from unemployment to employment and to nonparticipation. The model is based on continuous-time job search theory (see e.g. Mortensen, 1986). The model extends the non-stationary framework of Van den Berg (1990a) by allowing for endogenous nonparticipation, i.e. during unemployment individuals permanently have the option of exiting to nonparticipation. We allow both the rate at which job offers arrive and the wage offer distribution to change over the elapsed duration of unemployment. After presenting the outline of the model, we briefly discuss identification. We end this section with the parameterization.

Consider an unemployed worker being active on a labor market with 3 distinct labor market states: unemployment, employment and nonparticipation. Let t denote time, with $t = 0$ the moment of entering unemployment. During unemployment individuals receive job offers according to a Poisson process with an arrival rate $\lambda_t \geq 0$. A job offer is characterized by the wage w associated to the job offer. These wages are independent realizations from the wage offer distribution with continuous distribution function $F_t(w)$ (with finite mean, $\mathrm{E}_F[w|t] < \infty$). At the moment a job is offered, an individual has to decide immediately to accept this job or to reject it and continue searching. We do not allow for the possibility to reconsider job offers at a later stage. Once an unemployed worker decides to accept a job offer, he keeps the job forever at the same wage. We thus exclude on-the-job search and the possibility of losing a job. The instantaneous utility of being employed equals the wage received by the worker. Individuals discount future utility at the common subjective rate $\rho \geq 0$.

Like Calmfors and Lang (1995), we consider nonparticipation as an absorbing state. An individual in this state receives a (time-invariant) instantaneous utility, u_{np}. The model could be generalized by allowing the utility of being nonparticipant to be varying over the unemployment duration. This is shown in the formal exposition in the Appendix. While being unemployed, individuals receive a constant benefit level $b \geq 0$, which is also the instantaneous utility of being unemployed.

Consider the case in which the horizon is infinite and the individual has perfect foresight, i.e. he knows for each $t \geq 0$ his values of λ_t and $F_t(w)$, and anticipates future changes in these parameters. Unemployed workers do not know in advance when job offers arrive and what the associated wages are. Individuals maximize their expected present value of future utility. This implies for unemployed workers that they move to nonparticipation once the expected present value of continuing searching for jobs is lower than the utility of being nonparticipant. The essential point is that the transition to nonparticipation is an individual choice rather than a risk faced by the unemployed worker. Let R_t denote the expected present value of searching if the elapsed unemployment duration equals t. The Bellman's equation for R_t satisfies

$$R_t \;=\; \max\left\{\frac{u_{np}}{\rho}, b\triangle t + \lambda_t\triangle t(1-\rho\triangle t)\mathrm{E}_{F_t}\left[\max\left\{\frac{w}{\rho} - R_{t+\triangle t}, 0\right\}\right]\right.$$
$$\left. + (1-\rho\triangle t)R_{t+\triangle t}\right\} \qquad (1)$$

¿From the Bellman's equation it is easy to see that if the expected present value of continuing searching at t exceeds the present value of nonparticipation ($R_t > \frac{u_{np}}{\rho}$), the unemployed worker continues to search for a job and once he receives a job offer it is optimal to accept it if the associated wage exceeds ρR_t, which is denoted as the reservation wage at t, ϕ_t. The optimal strategy of the unemployed worker is therefore characterized by a sequence of reservation wages (ϕ_t, $\forall t \geq 0$) and the maximum duration of staying unemployed before becoming nonparticipant ($\bar{t}$). The unemployed worker accepts the first job offer which has a wage above the reservation wage or he moves to nonparticipation if his unemployment duration exceeds the maximum length of unemployment. The decision problem of the unemployed worker thus reduces to determining the sequence of reservation wages. Given the optimal value of $\bar{t}$, the optimal path of the reservation wage for $0 \leq t \leq \bar{t}$ is given by

$$\frac{\partial \phi_t}{\partial t} = \rho\phi_t - \rho b - \lambda_t \int_{\phi_t}^{\infty} (w - \phi_t)\, dF_t(w) \qquad (2)$$

The differential equation (2) for the reservation path, also given in Mortensen (1986) and Van den Berg (1990a), is well known and has been analyzed extensively. The way to ascertain the optimal time to go into nonparticipation is to solve the maximization problem

$$\max_{\bar{t}\geq 0} \phi_0 | (\phi_{\bar{t}} = u_{np})$$

¿From this equation it follows that $\bar{t}$ has to be found by solving ϕ_0 for all possible values of $\bar{t}$, which is time-consuming because the calculation of ϕ_0 requires the solution to the entire path of reservation wages. The Appendix gives algorithm for finding $\bar{t}$ and provides conditions under which $\bar{t}$ is unique. It might be clear that unemployed workers only enter nonparticipation if the reservation wage decreases over some range of unemployment duration.

The transition rate from unemployment to employment equals

$$\theta(t) = \lambda_t(1 - F_t(\phi_t)) \qquad \forall t \leq \bar{t} \qquad (3)$$

and does not exist for $t > \bar{t}$. Let t be the realized duration when leaving to employment. The conditional density function of t can be written as

$$f(t) = \theta(t) \exp\left(-\int_0^t \theta(s)ds\right) \qquad \forall t \leq \bar{t}$$

In Figure 1 we show how the reservation wage path changes when taking nonparticipation into account as a permanent option for the unemployed worker. In the beginning of the unemployment spell the reservation wage allowing for nonparticipation lies slightly above the reservation wage without nonparticipation. The difference between these reservation wage paths increases during the spell of unemployment. Since the reservation wage path is affected by allowing for nonparticipation also the acceptance probability of a job offer and therefore also the transition rates to work and the average post-unemployment wages change. In particular, the acceptance probability of a job offer decreases which in turn decreases the re-employment rate.

2.2 Some remarks on the identification

In this subsection, we briefly discuss the identification of the structural elements of the model. These structural elements are the wage offer distribution $F_t(w)$, the job offer arrival rate λ_t, the discount rate ρ, and the instantaneous utility of nonparticipation u_{np}. The identification is very much in line with Flinn and Heckman (1982).

Let us for a moment assume that we observe the unemployed worker's reservation wage path. From the accepted post-unemployment wages after an unemployment spell of t periods we can identify the wage offer distribution above the reservation wage $F_t(w|w > \phi_t)$. It is well known from Flinn and Heckman (1982) that the tail of the wage offer distribution below the reservation wage cannot be identified.

The re-employment hazard after unemployment duration t is given by $\lambda_t(1 - F_t(\phi_t))$. This hazard can be identified from the observed unemployment durations. However, it implies that the job offer arrival rate can only be identified up to a normalization. A high job offer arrival rate associated with a wage offer distribution that has some mass close to 0 cannot be distinguished from a low job offer arrival rate. To establish identification we will assume that the shape of the wage offer distribution is known up to a time-varying mean and an unknown standard deviation.

The instantaneous utility of being nonparticipant u_{np} can be identified from the length of the unemployment spell until entering nonparticipation. If the unemployed worker enters nonparticipation after an unemployment duration of t periods, the instantaneous utility of nonparticipation equals the reservation wage after t periods, i.e. $u_{np} = \phi_t$.

Finally, the discount rate ρ is identified from the observed reservation wage path. In the differential equation (2) describing the optimal reservation wage path all elements except for ρ are identified or observed. So solving the optimal reservation wage path identifies the discount rate ρ. In particular, to identify ρ it is only necessary to observe the reservation wage ϕ_t and the first derivative of the reservation wage with respect to the unemployment duration $\partial\phi_t/\partial t$ at one particular unemployment duration.[3]

So far, we have supposed that the complete reservation wage path of the unemployed worker is observed. However, the data do not provide the complete reservation wage path, but only observations of the reservation wage if the worker was unemployed at the moment of the interview. Since the moment of the interview is unrelated to the start of the unemployment spell, for different individuals we observe different parts of the reservation wage path. However, Flinn and Heckman (1982) stress that even if reservation wages are unobserved, similar identification results can be derived. The identification then hinges on the fact that the minimum of the accepted wages after unemployment duration t equals the reservation wage after unemployment duration t.

2.3 Parameterization

In this subsection, we discuss the parameterization of the unknown structural elements introduced in the previous subsections. We denote the vector of observed individual characteristics by x.

The structural model can be estimated in continuous time if it has a closed form solution for the reservation wage path. Only for very restricted specifications of the wage offer distribution such a closed form solution exists. Since we want to avoid an overly restrictive parametric specification, we maintain flexibility by estimating the model as a discrete time dynamic programming model (see e.g. Eckstein and Wolpin, 1989). In the data, the unit of time equals one month. However, we allow each month to include four periods of search of equal time length in which an individual can receive at most one job offer.[4]

[3] In practice, observing the reservation wage with respect to the unemployment duration $\partial\phi_t/\partial t$ would imply observing the reservation wages at two unemployment durations close to each other.

[4] The search unit of time should be sufficiently small, such that unemployed workers receive at most one job offer within the unit period. We started with a unit of time equal to one

Since we estimate the model in discrete time instead of continuous time, we specify a job offer arrival probability instead of a rate. We allow this probability to vary over the duration of unemployment and over both observed and unobserved individuals characteristics. The unobserved characteristics are denoted by v_λ. We use a logistic distribution to specify the job offer arrival probability

$$\lambda_t(x, v_\lambda) = \frac{\exp(\lambda(t) + x'\beta_\lambda + v_\lambda)}{1 + \exp(\lambda(t) + x'\beta_\lambda + v_\lambda)}$$

For the wage offer distribution we use a lognormal distribution function with parameters $\mu_t(x, v_\mu)$ and σ_w^2, where v_μ is the unobserved heterogeneity term in the wage offer distribution. The lognormal distribution function is the most common used specification for the wage offer distribution in the empirical literature on this type of job search model (e.g. Narendranathan and Nickell, 1985; and Wolpin, 1987). It is convenient since it generates relatively easy forms for the reservation wage path. Since σ_w^2 is assumed to be constant between individuals and over the duration of unemployment, all changes in the wage offer distribution come from changes in $\mu_t(x)$. For this parameter we choose an additive specification

$$\mu_t(x, v_\mu) = \mu(t) + x'\beta_\mu + v_\mu$$

If we would only observe one unemployment spell for each individual, we would not be able to identify the distribution of v_μ. However, for a number of individuals we observe more than one unemployment spell, whcih allows us to actually include an observed heterogeneity term in the wage offer distribution.

For both types of duration dependence $\lambda(t)$ and $\mu(t)$ we choose relatively flexible specifications,

$$\lambda(t) = \sum_{i=0,3,6,\ldots} \left(\frac{3-(t-i)}{3}\lambda_i + \frac{t-i}{3}\lambda_{i+3} \right) I(i \le t < i+3)$$

and

$$\mu(t) = \sum_{i=0,3,6,\ldots} \left(\frac{3-(t-i)}{3}\mu_i + \frac{t-i}{3}\mu_{i+3} \right) I(i \le t < i+3)$$

with $I(\cdot)$ being the indicator function. The functional forms imply that we specify the value of λ_t and μ_t at the beginning of each three months period and assume

month. Since the 'monthly' estimated job offer arrival probabilities were close to 1 for some unemployed, we shortened the unit of time. An estimated job offer arrival probability close to 1 implies a non-negligible probability of multiple job-offers within a time period, which contradicts the model and could therefore bias the other estimation results.

linear changes during the quarter. For the distribution of unobserved heterogeneity in the job offer arrival rate and the wage offer distribution, we choose a bivariate discrete distribution with two unrestricted mass-point locations for each term. Let v^a_λ, v^b_λ, v^a_μ and v^b_μ denote the points of support of v_λ and v_μ, respectively. The associated probabilities are denoted as follows:

$$\begin{array}{ll} \Pr(v_\lambda = v^a_\lambda, v_\mu = v^a_\mu) = p_{aa} & \Pr(v_\lambda = v^b_\lambda, v_\mu = v^a_\mu) = p_{ba} \\ \Pr(v_\lambda = v^a_\lambda, v_\mu = v^b_\mu) = p_{ab} & \Pr(v_\lambda = v^b_\lambda, v_\mu = v^b_\mu) = p_{bb} \end{array}$$

with $0 \leq p_i \leq 1$ for $i = aa, \ldots, bb$, and $p_{bb} = 1 - p_{aa} - p_{ab} - p_{ba}$.

Finally, as mentioned in the previous subsection, the instantaneous utility of being nonparticipant is considered to be constant during the unemployment spell. However, we do allow this to depend on both observed (x) and unobserved (v) individual characteristics and on the level of unemployment benefits

$$u_{np}(b, x, v) = \exp([1 \;\; \log(b) \;\; x']\beta_{np} + v_{np})$$

where v_{np} is distributed according to a normal distribution function with mean 0 and variance σ^2_{np}.

To be able to compute the reservation wage path backwards we impose that after some fixed unemployment duration T, both the job offer arrival rate $\lambda(t)$ and the wage offer distribution function $F_t(w)$ remain constant. Basically, this serves as a initial condition establishing a unique solution of the reservation wage path. We take this fixed duration T to be two years.

3 Data and institutions

In this section we first briefly discuss some institutional aspects in Germany during the observation period, which is from January 1989 until December 1995, and we provide some aggregated statistics. Then we give an overview of the data used in the empirical analyses.

3.1 Institutions and background information

The German unemployment benefits system contains three types of unemployment compensation schemes. If a worker involuntarily loses his job he is either entitled to Unemployment Insurance (UI) benefits (*Arbeitslosengeld*) or to Unemployment Assistance (UA) benefits (*Arbeitslosenhilfe*). The worker receives UI benefits if the worker was employed during 360 days out of the past 3 years. The UI benefits are of limited duration, with the entitlement period depending on the worker's age and his employment history. The requirement for receiving UI benefits are being registered at the labor office and being available and actively searching for a job. After the UI benefits period expires the worker is eligible for UA benefits, which are of unlimited duration.

The UI benefits level is 68% of the previous income (net of taxes) for individuals with dependent children and 60% for unemployed workers without dependent children. The levels of UA benefits are 58% and 53% of the previous earnings respectively. For both UI en UA there is a minimum and maximum level of benefits. The drop in benefits level is thus relatively small when the eligibility period of UI benefits expires. Compared to other countries, such as the US where unemployment benefits eventually run out, in Germany the level of unemployment benefits is very static.[5]

Individuals who do not qualify for UI or UA benefits are eligible for collecting welfare benefits (*Sozialhilfe*). For example, unemployed individuals, whose previous state was full-time education, receive these benefits. Welfare benefits are means-tested and of unlimited duration. The level of welfare benefits depends on the household situation and the age of the dependent children. The welfare benefits do not require the unemployed worker to actively search for work.

Ljungqvist and Sargent (1998) present some (empirical) net replacement rates for Germany. In 1994 the net replacement rate for single unemployed workers was 66% and after the first year of unemployment it decreased to 63%, where it remained constant until the fifth year of unemployment. For unemployed workers with a dependent spouse the replacement rates were 74% during the first year of unemployment and 72% from the second year onwards.

In Figure 2 we present the unemployment rate in Germany during the observation period. We also show the unemployment rates stratified to former East-Germany and West-Germany. Reliable unemployment rates for former East-Germany from before 1991 are not available. The figure shows that during the observation period, the unemployment rate in East-Germany is almost twice that of West-Germany. The level of the unemployment rate remains relatively constant during the observation period. Only in 1991 was the unemployment rate in East-Germany lower than during the period thereafter. However, this might have to do with the start of collecting aggregated unemployment figures for East-Germany.

Between 1990 and 1994, in Germany only 0.5% of the workers between 15 and 44 years old were collecting some type of disability benefits. This percentage is particularly low compared to other countries such as the U.S., the U.K., Sweden,

[5]See Hunt (1995) for a more extensive discussion of the net replacement rates.

where in the same period around 2.5% of the workers received disability benefits (see Buddelmeyer, 2001). For older workers in Germany, the percentage of workers collecting disability benefits increases dramatically. This is in accordance with the German labor force participation. For both men and women the labor force participation is high compared to other countries for workers until age 50 and than it decreases rapidly (see OECD, Employment Outlook, various years). On the other hand, part-time work and short-working hours are less common in Germany than in most other countries.

Another interesting indicator is the percentage of unemployed workers who have been long-term unemployed. Ljungqvist and Sargent (1998) report that in Germany the percentages remain very constant during our observation period. In 1989, 66.7% of the stock of unemployed workers was already unemployed for more than 6 months and 49.0% for more than a year. These percentages where slightly lower in 1995, respectively 65.4% of the unemployed workers was unemployed for more than 6 months and 48.3% longer than a year. Machin and Manning (1999) compare for 1995 the unemployment rates and percentages of long-term unemployed workers between different countries. They show that, compared to the unemployment rate in Germany, the percentage of long-term unemployed workers is relatively high. This implies that in Germany unemployment is more persistent than in other countries.[6] Finally, Blanchard and Wolfers (2000) argue, based on measures summarized in the OECD Employment Outlook (1999), that in Germany employment protection has been roughly stable since the early 1970s.

3.2 Data

The micro data we use for the empirical analyses are from the German Socio-Economic Panel (GSOEP), which is a longitudinal panel survey. An extensive discussion on the construction of this database is given in Wagner, Burkhauser and Behringer (1993). The GSOEP started in 1984 in West-Germany. East-German households have been included since 1990. The total number of individuals included in the database is about 20,000.

Yearly, all members of the households older than 16 years are interviewed. Those who leave the household stay in the panel as long as they remain in Germany. People who join households with original members of the panel are also interviewed as long as they stay in such a household. The attrition from the panel is low and is compensated by household members reaching the age of 16

[6] Both Ljungqvist and Sargent (1998) and Machin and Manning (1999) base their figures on the OECD Economic Outlook (1991, 1995).

who thereby enter the panel.

The GSOEP contains extensive information on the labor market behavior of individual respondents in the year preceding the moment of interview. In particular, respondents are asked to report their labor market states for each month in the previous year. The respondent can choose between 12 states:

1. Full-time employment
2. Short-working hours
3. Part-time employment
4. Vocational training
5. Unemployment
6. Retirement
7. Maternity leave
8. School and college
9. Military and civil service
10. Housewife or househusband
11. Second job
12. Other

We consider states 1-3, 9 and 11 as employment states; states 7, 10 and 12 are the nonparticipation states and state 5 is the unemployment state. State 6 does not arise in our final sample (see below). By comparing individual labor market states in consecutive months, it is possible to construct spells of unemployment. These spells always consist of an integer number of calendar months. If a spell of unemployment is interrupted by a short spell of training (state 4 or 8), we ignore these transitions and consider the interrupted spell as a single unemployment spell.[7] Unemployment is self-reported and is hence not necessarily equal to

[7]During the last decades a large variety of publicly funded training and schooling programs for unemployed workers has become available in Germany (see e.g. Hujer, Maurer and Wellner, 1999). Because the main objective of these programs is stimulating re-employment, an unemployed worker who enters the state of training or schooling can continue searching for jobs. We thus follow Pannenberg and Schwarze (1996), who investigated measures of labor market tightness, by considering this training as a continuation of unemployment.

being registered at an unemployment agency. Similarly, nonparticipants may be registered at an unemployment agency and receive unemployment benefits.

Because we are mainly interested in reasons for leaving the labor force other than retirement, we restrict the data set to the 2044 individuals between 17 and 45 years old. We exclude 106 individuals for which 'region of residence', 'gender' or 'years of education' is missing. For the remaining 1938 individuals, we construct a flow sample by considering spells of unemployment starting between January 1989 and December 1995. The individuals in our data set experienced 2649 spells of unemployment. As mentioned in Subsection 2.3, we only model the job search behavior during the first 2 years of unemployment. Therefore, we artificially right-censor unemployment spells after 2 years. Of the unemployment spells 68% ends because the unemployed workers became employed and 9% of the spells of unemployment ends in nonparticipation. The remaining 23% of the unemployment spells is right-censored. Because most right-censoring occurs at the end of the observation period (the month of the last interview), we treat it as exogenous.

An important issue is the level of unemployment benefits. Recall that the effect of benefit profiles on reservation wage paths is very large. For 952 spells we do not observe the benefits level. It seems very unlikely that these individuals have no income at all. Most of these spells are spells with a short unemployment duration: 490 of those spells had an unemployment duration less than 3 months and 878 less than 1 year. Excluding these spells from the data set would cause a bias toward long-term unemployment spells and would thus affect duration dependence. Therefore, we impute for the spells with a missing benefits level, a level of unemployment benefits. We stratify the data set into 4 subsamples based on the observed unemployment duration, a subsample with spells shorter than 3 months, a subsample with spells between 3 and 6 months, a subsample with spells between 6 months and 1 year and a subsample with spells over 1 year. For each of these subsamples we regressed the log benefits level on a set of individual characteristics including the previous wage. If the previous wage was unobserved we replaced it with the post-unemployment wage. If the benefits level is missing we impute it with a draw from the log-linear regression.

The personal characteristics of the respondents are recorded at the beginning of the spell of unemployment and are considered time-invariant. The set of characteristics include the age in years, an indicator function for having children, years of education, being female, having a non-German nationality, and the region in which the unemployed worker lives.[8] Furthermore, for most of the individuals we

[8]We distinguish 5 regions in Germany; South-Germany (Bavaria and Baden-Wurttemberg),

observe the marital status. For individuals for whom this is missing, we add a dummy variable that indicates missing. We also observe the spell-specific characteristic "labor market state before inflow into unemployment". The variable indicates if the individual entered unemployment out of employment or from out of the labor force. The latter type are mainly new entrants on the labor market. Finally, we have observations on individually reported reservation wages in the years 1992, 1993 and 1994. Because the reservation wage was only asked if the worker was unemployed at the date of the interview, we only observe the individual reservation wage for 14% of the spells. Hence the individual reservation wage is not always registered at the beginning of a spell. We return to this issue in Subsection 5.1.

Table 1 provides some statistics of the data set. Females, individuals with children, married individuals, residents of North-Germany and individuals who entered unemployment after having been out of the labor force are more likely to exit unemployment to nonparticipation. However, most striking is the difference in average individual reservation wage corresponding to unemployment spells exiting to employment and to nonparticipation. The average individual reservation wage is around 300 D-Mark lower if the spell ended in nonparticipation. This suggests that unemployed workers who enter nonparticipation have worse labor market prospects than other unemployed workers, which indicates a substantial level of selectivity for transitions to nonparticipation. Finally, the correlation between the unemployment duration and the post-unemployment rate is -0.15, indicating that individuals who have been unemployed for a relatively long period have a lower post-unemployment wage. This might either be caused by genuine duration dependence or by heterogeneity of unemployed workers.

Figure 3 shows how the (empirical) monthly exit probabilities to employment and nonparticipation change over the duration of unemployment. The probabilities of a transition to a job clearly decrease over the duration of unemployment. However, this type of figure is unable to distinguish between genuine duration dependence and (observed and unobserved) heterogeneity. We return to this issue in Section 4, where we perform some reduced-form analyses. The transition probabilities to nonparticipation increase during the first year of unemployment and thereafter decrease. In a stationary environment, such a pattern could not occur. In our non-stationary environment, a stronger decreasing reservation wage

Mid-(West-)Germany (Saarland, Rhineland-Palatinate and Hesse), North-Germany (Bremen, Hamburg, Lower Saxony, North Rhine Westphalia and Schleswig-Holstein), (former-)East-Germany (Brandenburg, Mecklenburg-West Pomerania, Saxony, Saxony-Anhalt and Turingia), and Berlin. The grouping is on the basis of average incomes.

at some time interval causes a peak in transition probabilities.

4 Reduced-form analyses

In this section we perform some reduced-form analyses. The results provide a benchmark for the predictions of the structural model. We estimate a competing risks model on the unemployment duration until exit to work or exit to nonparticipation. Furthermore, we estimate a wage equation which allows the wage in the first job after unemployment to depend on observed and unobserved individual characteristics and on the length of the previous spell of unemployment. In both models, the parameters are estimated using Maximum Likelihood.

4.1 Unemployment duration

We use the standard continuous-time hazard rate framework to model the transition rates from unemployment to work and to nonparticipation (see e.g. Lancaster, 1990; and Van den Berg, 2001). We assume that differences in transition rates can be characterized by observed characteristics x, unobserved characteristics v and the elapsed unemployment duration t itself. Also we assume both x and v to remain constant within an unemployment spell and v to be independent of x. In particular, we impose that v is constant within all unemployment spells of a given individual, i.e. we exploit that the data provide multiple unemployment spells for some workers.

The transition rates from unemployment to work and to nonparticipation at t conditional on x and v are denoted by $\theta_w(t|x, v_w)$ and $\theta_n(t|x, v_n)$. Both hazard rates are assumed to have the familiar Mixed Proportional Hazard (MPH) specification

$$\theta_d(t|x, v_d) = \lambda_d(t) \exp(x'\beta_d + v_d) \qquad d = w, n$$

in which $\lambda_d(t)$ represents individual (baseline) duration dependence.

It may be clear that we can only observe one transition, either to work or to nonparticipation, whichever occurs first. Let the indicator function $I(w)$ be 1 if a transition to work is observed and $I(n)$ is 1 if the unemployment spell is observed to exit to nonparticipation. We do not observe the exact date at which the spell of unemployment starts and ends, but only the month during which it started and ended. For simplicity, we assume that a spell of unemployment always starts at the beginning of a month. However, when computing the contributions to the likelihood function, we integrate over the month in which the spell ended. Conditional on the unobserved individual characteristics (v_w, v_n), the (conditional) contribution to the likelihood function of an unemployment spell with length of $t = 1, 2, \ldots$ months is

$$\begin{aligned}\ell(t, I(w), I(n); x, v_w, v_n) &= \int_{t-1}^{t} \theta_w(s|x, v_w)^{I(w)} \theta_n(s|x, v_n)^{I(n)} \\ &\qquad \exp\left(- \int_0^s \theta_w(z|x, v_w) + \theta_n(z|x, v_n)\, dz \right) ds\end{aligned}$$

The use of a flow sample of unemployment spells implies that we do not have any initial conditions problems. The right-censoring in the data is exogenous and is therefore solved in a straightforward manner within the hazard rate framework. In case of independence of v_w and v_n, we would have a standard duration model for both the duration until exit to work and to nonparticipation in which right-censoring of the duration until work due to a transition in nonparticipation can be treated as exogenous and vise versa. However, if v_w and v_n are not independent, inference has to be based on the joint distribution of $(t, I(w), I(n))|x$.

Let $G(v_w, v_n)$ be the joint distribution function of the unobserved characteristics (v_w, v_n). Let t_{ij} denote the duration (in integer months) of the j^{th} spell of unemployment of individual i ($i = 1, \ldots, N$ and $j = 1, \ldots M_i$). The (unconditional) individual contribution to the likelihood function equals

$$\ell_i = \int_{v_w} \int_{v_n} \prod_{j=1}^{M_i} \ell(t_{ij}, I_{ij}(w), I_{ij}(n); x_{ij}, v_w, v_n)\, dG(v_w, v_n)$$

The log-likelihood function, which we optimize to obtain the parameter estimates equals

$$\log \mathcal{L} = \sum_{i=1}^{N} \log(\ell_i)$$

Because we observe multiple spells, the competing risks model is identified under very weak conditions. From Heckman and Honoré (1989) it follows that under general conditions the whole model is identified from the data corresponding to the competing risks part.

For the duration dependence functions and the bivariate unobserved heterogeneity distribution we take the most flexible specifications used to date. We take both $\lambda_w(t)$ and $\lambda_n(t)$ to have a piecewise constant specification,

$$\lambda_d(k) = \exp\left(\sum_{k=1,2,\ldots} \lambda_{dk} I_k(t) \right) \qquad d = w, e$$

where k is a subscript for time intervals and $I_k(t)$ are time-varying dummy variables that are one in consecutive time intervals. With an increasing number of time intervals any duration dependence pattern can be approximated arbitrarily closely, which allows us to avoid duration dependence specifications with only one parameter (like a Weibull specification) that are well known to be overly restrictive (see e.g. Lancaster, 1990).

We take the joint distribution of the unobserved heterogeneity terms v_w and v_n to be bivariate discrete with unrestricted mass-point locations for each term. We found that it is optimal to have three points of support for both the hazard rate to work and to nonparticipation. Adding additional points of support did not improve the fit of the model. Let v_w^a, v_w^b, v_w^c, v_n^a, v_n^b and v_n^c denote the points of support of v_w and v_n, respectively. The associated probabilities are denoted as follows:

$$\begin{array}{lll}
\Pr(v_w = v_w^a, v_n = v_n^a) = p_{aa} & \Pr(v_w = v_w^a, v_n = v_n^b) = p_{ab} & \Pr(v_w = v_w^a, v_n = v_n^c) = p_{ac} \\
\Pr(v_w = v_w^b, v_n = v_n^a) = p_{ba} & \Pr(v_w = v_w^b, v_n = v_n^b) = p_{bb} & \Pr(v_w = v_w^b, v_n = v_n^c) = p_{bc} \\
\Pr(v_w = v_w^c, v_n = v_n^a) = p_{ca} & \Pr(v_w = v_w^c, v_n = v_n^b) = p_{cb} & \Pr(v_w = v_w^c, v_n = v_n^c) = p_{cc}
\end{array}$$

with $0 \leq p_i \leq 1$ for $i = aa, ab, ac, ba, bb, bc, ca, cb, cc$, and $p_{cc} = 1 - p_{aa} - p_{ab} - p_{ac} - p_{ba} - p_{bb} - p_{bc} - p_{ca} - p_{cb}$.

The estimation results are given in Table 2. Although we defined 9 possible combinations of mass-point for the distribution of the unobserved heterogeneity, we only found 3 combinations which have a positive probability. In particular, the estimation results indicate a negative correlation between the unobserved heterogeneity term in the employment hazard and nonparticipation hazard, implying that individuals who have a low probability of finding employment are more likely to enter nonparticipation.

Females and individuals who were out of the labor force before entering unemployment have both a lower exit rate to employment and a higher exit rate to nonparticipation than males and previously employed individuals, respectively. Years of education and nationality only significantly affect the exit rate to work, while age, having children, and marital status only have a significant effect on the rate at which individuals leave the labor force. The corresponding parameter estimates have the expected sign, although it might be surprising that young unemployed workers are more likely to enter nonparticipation than old unemployed workers. However, recall that we only consider individuals under the age of 45, so these individual are not eligible for early retirement yet. Unemployed workers living in East-Germany (and to a lesser degree in Berlin) have both lower exit rates to work and to nonparticipation than unemployed workers living in West-Germany.

Is is not possible to assign a structural interpretation to these observed covariate effects. Recall that both the exit rate to employment and to nonparticipation are affected by all structural elements in the model. A higher job offer arrival rate increases the exit rate to work and decreases the transition rate into nonparticipation. So one might argue that males should have higher job offer arrival rates than females. However, also a lower instantaneous utility of nonparticipation increases the exit rate to work (as is decreases the reservation wage) and increases the transition rate into nonparticipation. It might therefore also be the case that females have a higher utility of being nonparticipant than males. These possibilities can be disentangled with the structural model.

The level of unemployment benefits only has a significant impact on the reemployment rate. Unemployed workers who receive high benefits have a lower exit rate to work. This is in accordance with the predictions from the structural model, where higher benefits increase the reservation wage and hence decrease the transition rate to work. If the utility of nonparticipation would not depend on the benefits level, a high benefits level would decrease the transition rate into nonparticipation. The level of unemployment benefits has a small and insignificant negative impact on the transition rate to work.

The structural model predicts that reservation wages should be decreasing during the unemployment spell, because otherwise transitions into nonparticipation would not have been observed. A decreasing reservation wage path can be associated with both negative duration dependence in the job offer arrival rate and in the wage offer distribution. Both types of negative duration dependence also cause the exit rate to employment to decrease, which is confirmed by the estimation results of the baseline hazard in the exit rate to work. The shape of the exit rate to nonparticipation depends on how the utility of nonparticipation is distributed within the population of unemployed workers (and on the exact reservation wage path). The duration dependence pattern in the transition rate to nonparticipation is hump-shaped, which implies that during the first year of unemployment the exit rate into nonparticipation increases and afterwards it decreases again.

The MPH structure imposes that the baseline hazard is the same for all individuals. From the structural model it follows that this is not the case, because identical unemployed workers should enter nonparticipation after exactly the same unemployment duration. This means that the baseline hazard should be dependent on individual characteristics. Parameter estimates of the reduced-form model might therefore be biased.

4.2 Post-unemployment wages

In this subsection we investigate the post-unemployment wage. We estimate a wage equation in which we allow the wage after the j-th unemployment spell of individual i to depend on observed characteristics x_{ij} (including the level of unemployment benefits, the previous labor market state, and the length of the past unemployment spell), and an individual-specific effect v_i. In particular, we specify the logarithm of the wage as a linear function of these explanatory variables,

$$\log(w_{ij}) = \alpha + x'_{ij}\beta + v_i + \varepsilon_{ij}$$

Furthermore, we assume that the individual-specific effects v_i are random effects distributed according to a normal distribution with mean 0 and variance σ_v^2. Also ε_{ij} is assumed to follow a normal distribution with mean 0 and variance σ_ε^2. This equation is estimated with Maximum Likelihood.

Table 3 shows the estimation results of the random effects wage equation. The parameter estimates show that age, gender, years of education, the level of unemployment benefits, region and the previous labor market state are the most important individual characteristics. The post-unemployment wages of people in (former) East-Germany is lower than in other parts of Germany. The same holds for younger workers, females, and low-educated individuals who receive a lower post-unemployment wage. Individuals receiving more unemployment benefits have, on average, higher post-unemployment wages. The structural model predicts that unemployed workers with higher benefits have a higher reservation wage. Therefore, they are more selective in the job offers they accepted and they receive higher wages. Also, individuals who were employed previous to unemployment receive higher expected post-unemployment wages compared to individuals who entered unemployment after having been out of the labor force. This latter group consists mainly of individuals who enter the labor force for the first time and thus have no work experience. The effect of previous benefits may reflect the effect of having previous work experience.

¿From the estimation result of this reduced-form analysis it remains unclear what drives the differences in average wages between groups of individuals. On the one hand, the job offer arrival rates between individuals might differ, which means that unemployed workers with a low job offer arrival rate have lower reservation wages and hence accept lower wages. On the other hand, some workers might have less skills and therefore receive lower wage offers. To distinguish between these competing hypotheses, the structural model is needed.

The effect of the duration of the past unemployment spell on the post-unemployment wages are closely related to the predictions of the structural model. The length of the past unemployment spell has a negative effect on wages. One year of unemployment decreases post-unemployment wages by almost 7%. This implies that the structural model is non-stationary. Again the structural model is needed to distinguish between changes in the wage offer distribution and changes in the job offer arrival rate.

Finally, the reduced-form wage equation does not take explicit account of the selectivity of observed wages, i.e. we only consider observations for which we actually observe a post-unemployment wage. However, some people might still have been unemployed or might have entered nonparticipation at the end of the observation period. It is well known that this endogenous selectivity may cause biased parameter estimates. The structural model does take account of this selectivity by explicitly modeling the process through which unemployed workers enter employment.

5 Estimation of the structural model

In this section we estimate the structural model presented in Section 2. Before discussing the parameter estimates, we consider some preliminary issues and briefly discuss the estimation method. After presenting the results of the structural model, we compare them with those of the reduced-form analysis and we discuss the policy implications of the estimated structural model.

5.1 Some preliminary issues

For some individuals we observe the reservation wage at some moment during their unemployment spell and/or the accepted post-unemployment wage. We assume that both are observed with a measurement error. Let $\tilde{\phi}_t$ be the observed reservation wage after an unemployment duration of t periods and ϕ_t denotes the true reservation wage at this moment, i.e. the reservation wage predicted by the model. We assume that these are related by $\log(\tilde{\phi}_t) = \log(\phi_t) + \varepsilon_\phi$, where ε_ϕ is normally distributed with mean 0 and variance η_ϕ^2. Similarly, we assume that the observed accepted wage $\tilde{w}_t$ and the true post-unemployment wage w_t are related according to $\log(\tilde{w}_t) = \log(w_t) + \varepsilon_w$. In this latter case, we do not know the true accepted wage but only know that it is distributed according to the truncated lognormal distribution function $F_t(w|w > \phi_t)$. The observed wage distribution is thus the convolution of a lognormal and truncated lognormal distribution function. This convolution has a relatively simple closed-form solution. Both η_ϕ^2 and η_w^2 can be interpreted as measures of the goodness of fit of the model. If the estimated values of these parameters are low, it implies that the model can describe the data well.

Another data issue concerns the instantaneous utility of nonparticipation. Since our data consist of a flow sample of unemployed workers, there is an initial condition problem. Individuals with a very high utility of nonparticipation would not have entered unemployment, but would have become nonparticipant immediately.[9] The distribution function of the unobserved component in the instantaneous utility of nonparticipation v is thus truncated from above. It is clear that we have to correct for this by considering the truncated distribution function of v instead of the complete distribution function in the computation of the likelihood function. Let $G(v)$ be the cumulative normal distribution function of v in the whole population, this implies that in our data v is distributed according to $G(v|u_{np}(v) < \phi_0)$.

In the remainder of this subsection we provide the individual contributions to the log-likelihood function. We can distinguish three types of unemployment spells, (*i*) a spell ending in nonparticipation, (*ii*) a right-censored spell, and (*iii*) a spell ending in employment. For each of these spells we briefly discuss the likelihood contributions. First, we introduce some additional notation. Let d_t be a dummy variable which indicates if the reservation wage after t periods is observed. As mentioned above, reservation wages are observed with a measurement error, where $h_\phi(\tilde{\phi}_t|\phi_t)$ is the lognormal density function for the discrepancy between the observed and true reservation wage. Similarly, we have the lognormal density function $h_w(\tilde{w}_t|w_t)$ for the observed wage. In Subsection 2.1 we only define ϕ_t until the moment of entering nonparticipation, $t \leq \bar{t}$ (see equation (2)). Since nonparticipants do not receive job offers, the reservation wage is undefined for $t > \bar{t}$. For ease of presentation we define $\phi_t = u_{np}$ for nonparticipants ($t > \bar{t}$). The values of λ_s and F_s depend on the unobserved heterogeneity v_λ and v_μ. Consequently also the reservation wage path ϕ_t depeds on these mass points. Below we suppress this dependence. However, for each possible set of mass points for v_λ and v_μ we compute the likelihood contributions, which are in the likelihood function weighted by their probability mass.

First, we focus on unemployment spells which end in nonparticipation. Consider an unemployed worker who is observed entering nonparticipation during

[9]Including individuals who entered nonparticipation without some period of unemployment in the analysis is unattractive. This group might for example also include formerly employed workers who became disabled. Since this is a risk rather than a choice, these type of transitions do not fit our model.

his t^{th} period of unemployment. The instantaneous utility of nonparticipation of this individual is below ϕ_{t-1} and is by definition equal to ϕ_t. Furthermore, this individual did not find a job during this period, either because he did not receive any job offers, or because he rejected all jobs offered. The contribution to the likelihood function of such an unemployment spell equals

$$\int_{-\infty}^{\infty} I\left(\phi_t \leq u_{np}(v) < \phi_{t-1}\right) \left(\prod_{s=1}^{t} \left(1 - \lambda_s \left(1 - F_s(\phi_s)\right)\right)\right) \left(\prod_{s=0}^{t-1} h_\phi(\tilde{\phi}_s|\phi_s)^{d_s}\right) dG(v|u_{np}(v) < \phi_0)$$

where $I(\cdot)$ is the indicator function taking the value 1 if the expression in parentheses is true and 0 if not. As mentioned earlier, ϕ_t depends on the instantaneous utility of nonparticipation u_{np} and thus also on v.

The likelihood contribution of a right-censored unemployment spell is similar to that of a spell ending in nonparticipation. The difference is that there is no lower bound for the instantaneous utility of nonparticipation. The contribution to the likelihood function is thus

$$\int_{-\infty}^{\infty} I\left(u_{np}(v) < \phi_t\right) \left(\prod_{s=1}^{t} \left(1 - \lambda_s \left(1 - F_s(\phi_s)\right)\right)\right) \left(\prod_{s=0}^{t-1} h_\phi(\tilde{\phi}_s|\phi_s)^{d_s}\right) dG(v|u_{np}(v) < \phi_0)$$

Finally, the last type of unemployment spells are spells ending into employment after t periods. The individual did not find a suitable job and did not move to nonparticipation during the first $t-1$ periods, and accepted a job in the t^{th} period. In particular, the wage w associated with the job offer exceeds the reservation wage after t periods. The contribution to the likelihood function of such spell is

$$\int_{-\infty}^{\infty} I\left(u_{np}(v) < \phi_{t-1}\right) \left(\prod_{s=1}^{t-1} \left(1 - \lambda_s \left(1 - F_s(\phi_s)\right)\right)\right) \lambda_t(1 - F_t(\phi_t)) \int_{w \geq \phi_t} h_w(\tilde{w}_t|w) dF_t(w) \left(\prod_{s=0}^{t-1} h_\phi(\tilde{\phi}_s|\phi_s)^{d_s}\right) dG(v|u_{np}(v) < \phi_0)$$

As mentioned above the integral $\int_{w \geq \phi_t} h_w(\tilde{w}_t|w) dF_t(w)$ has a closed-form solution and therefore needs no further attention.

In the derivation of the likelihood contributions above we assumed the data are measured in the same unit of time as the model is specified. Recall that the unit of time in the model is a week, while the data are measured in months. Therefore, if a transition to nonparticipation or to employment is observed in a particular month, then we have to integrate over the weeks in this month. Similarly, if a post-unemployment wage is observed, we have to integrate over the possible lengths of the unemployment spell. These integrals are solved straightforwardly.

5.2 Estimation method

The contributions to the likelihood functions mentioned in the previous subsection do not have a closed-form solution. In particular, the reservation wage path depends on the instantaneous utility of nonparticipation, which is unobserved. The other elements of the likelihood function, such as the job acceptance probability, depend on the reservation wage. To overcome this problem, we use simulation methods to compute the individual contributions to the log-likelihood function. See Pakes (1986) for an early application of simulation estimation of dynamic programming models and Bloemen (1997) for an application to a structural job search model. In this subsection we briefly discuss the simulated maximum likelihood method we use to estimate the structural model.

In our framework it would be most straightforward to sample a fixed number of draws for v from $G(v|u_{np}(v) < \phi_0)$ and then to evaluate the likelihood function for each sampled value of v. However, for many values of v the likelihood contributions are 0, because the argument within the indicator function $I(\cdot)$ is not satisfied. This particular value of v could not have generated the observed data, since the predicted month of entering nonparticipation is inconsistent with the observed unemployment duration. For a finite number of draws, these type of acceptance/rejection-procedures generally generate log-likelihood functions that are discontinuous in the parameters. If the log-likelihood function is not smooth in the parameters, convenient gradient methods are not useful in the optimization procedure (see Stern, 1997).

Therefore, we use an alternative approach in line with Keane (1994). We rewrite

$$I\left(\phi_t \leq u_{np}(v) < \phi_{t-1}\right) dG(v|u_{np}(v) < \phi_0) =$$

$$\frac{\Pr_G[\phi_t \leq u_{np}(v) < \phi_{t-1}]}{\Pr_G[u_{np}(v) < \phi_0]} dG(v|\phi_t \leq u_{np}(v) < \phi_{t-1})$$

note that if $u_{np}(v) < \phi_{t-1}$, then also $u_{np}(v) < \phi_0$. Next we determine the support of v for which $\phi_t \leq u_{np}(v) < \phi_{t-1}$, and let this support be denoted by $(\underline{v}, \overline{v})$. Similarly we determine the support $(-\infty, v_0)$ on which the condition $u_{np}(v) < \phi_0$ holds.[10] Now, we can simply rewrite $\Pr_G[\phi_t \leq u_{np}(v) < \phi_{t-1}] = \Phi(\overline{v}/\sigma_{np}) - \Phi(\underline{v}/\sigma_{np})$ and $\Pr_G[u_{np}(v) < \phi_0] = \Phi(v_0/\sigma_{np})$, where $\Phi(\cdot)$ is the standard normal distribution function. Since we impose that v follows a normal distribution function, generating realization from $G(v|\phi_t \leq u_{np}(v) < \phi_{t-1})$ is not very complicated. If u is a draw from a uniform distribution, then

$$v = \sigma_{np}\Phi^{-1}\left(u(\Phi(\overline{v}/\sigma_{np}) - \Phi(\underline{v}/\sigma_{np})) + \Phi(\underline{v}/\sigma_{np})\right)$$

is a realization from $G(v|\phi_t \leq u_{np}(v) < \phi_{t-1})$ (see Stern, 1997). Before starting the optimization routine of the log-likelihood function we generate for each individual a set $u_1, \ldots, u_S$ from a uniform distribution function, which remains constant during the complete optimization.

In general, for a finite number of draws simulated maximum likelihood is inconsistent even if the number of observations tends to infinity. Gourieroux and Montfort (1996) show that simulated maximum likelihood is consistent if both the number of draws per observation S and the number of observations N go to infinity. If also $\sqrt{N}/S$ goes to 0, simulated maximum likelihood estimation is asymptotically equivalent to ordinary maximum likelihood estimation. We take S equal to 25. This is in line with the Monte Carlo evidence provided by Börsch-Supan and Hajivassiliou (1993), which shows that the bias in the parameter estimates is negligible if S exceeds 20.

5.3 Estimation results

The estimation results of the structural model are provided in Table 4. Recall that the structural model is specified such that the expected present value of continuing job search must be decreasing at the moment an unemployed worker enters nonparticipation. This imposes a restriction on the support of the parameters of the structural elements. We have checked if the parameter estimates are in the interior of the support of the parameters. Since this is the case, the presented standard errors are based on the regular asymptotic properties of maximum likelihood estimation.

The estimation results show significant negative duration dependence in the wage offer distribution. The wage offer distribution decreases slightly faster during the first 15 month after becoming unemployed than afterwards. Duration

[10] We use numerical methods to compute $\underline{v}$, $\overline{v}$ and v_0. This can easily be done using condition (5) of Theorem 1 provided in the Appendix.

dependence in the job offer arrival rate is not significant, the p-value of a Wald-test for joint significance equals 0.96.

The unobserved heterogeneity in the job offer arrival rate and the wage offer distribution turns out to be significant. Around 71% of all probability mass is assigned to individuals with high job offer arrival rates and on average higher wage offers. 18% of the probability mass is located at a low job offer arrival rate, but on average higher wage offer distributions. This indicates that unobserved heterogeneity is more important for explaining the job offer arrival rate than the wage offer distribution. About 9% of the probability mass is located to both a low job offer arrival rate and on average lower wage offers and less than 2% to a high job offer arrival rate, but on average lower wage offers.

Years of education and being female are the only explanatory variable that have a significant impact on both the probability of receiving job offers and the wage offer distribution. Higher-educated and male unemployed workers not only receive more job offers on average, but also get offered higher wages on average. The other important individual characteristics in the job offer arrival rate are having children and the labor market state before unemployment. Unemployed workers with children and those who worked previous to unemployment have higher probabilities of receiving a job offer while being unemployed. The latter probably indicates recent work experience. Age is an important variable in the wage offer distribution. Older individuals receive on average higher wage offers than younger individuals. This can be explained from a human capital perspective, where age is a proxy for total work experience. The utility of nonparticipation is significantly higher for women, individuals with children, unemployed workers who were also nonparticipant before entering unemployment and individuals with higher unemployment benefits. This latter finding probably reflect the fact that there is a positive relation between the level of the benefits in nonparticipation and in unemployment. Since unemployment and nonparticipation in the data is self-reported many of the individuals who stop searching for work continue receiving unemployment benefits.

There are no significant regional differences in the probability of receiving job offers. However, the wages associated to job offers in East-Germany are significantly lower than in former West-Germany, which causes transition rates to employment to be lower in East-Germany. This is a possible labor demand explanation for this. After the German re-unification, large investments were made in East-Germany. Mainly low-skilled (construction) jobs became available, which explains why job offer arrival rates are relatively large and associated wages are on average relatively low.

The estimated yearly discount rate is around 19%, which is high, although well within the range found in the empirical studies reviewed by Frederick, Loewenstein and O'Donogue (2002). It is well known that the estimated discount rate is sensitive to the specification of the utility function. Our utility function assumes that unemployed workers are risk neutral. If in fact the unemployed workers are risk averse, the discount rates pick up some of the risk aversion and therefore the estimated discount rates overestimates the true discount rate. Furthermore, we have assumed that employment and nonparticipation are absorbing states. This is obviously not always the case, which affects the estimated reservation wage.

5.4 Fit of the model and policy implications

As mentioned in Subsection 5.1, the estimated variance of the measurement errors in the observed wages and reservation wages give an indication of the goodness of fit of the model. In the data, the variance in the logarithm of wages is 0.16 and the estimated variance of the measurement error of wages, $\hat{\eta}_w^2$, is 0.067. This implies that the model explains 58% of the variation in the logarithm of observed wages, which is reasonable. The estimated variance in the measurement error of the reservation wages can be decomposed as $\hat{\eta}_\phi^2 = \mathrm{E}[\log(\tilde{\phi}) - \log(\phi)]^2 + \mathrm{var}(\log(\tilde{\phi}) - \log(\phi))$. The first component denotes the square-root of the bias, which equals 0.00012. The model structurally overestimates the reservation wages with around 1.1%. The second component is the unexplained variance, which equals 0.074. Since the variance in the logarithm of the reservation wages equals 0.13, our model can explain only 42% of the variance in logarithm of reservation wages. There is thus hardly any systematic bias in the reservation wages, but the fit is not extremely good.

Another indication for the goodness of fit of the model is comparing the estimation results with the estimation results of the reduced-form model discussed in Section 4 and the actual data. Table 5 provides the fractions of unemployment spells that end in work and in nonparticipation within a given number of months based on the actual data, the estimated competing risks model and the structural model.[11] The fit of the structural model seems reasonably good. In particular, the fraction of unemployment spells that exit into work is estimated properly. The structural model underestimates the number of transitions into nonparticipation. In this latter aspect the reduced-form model performs better than the structural model.

[11] We have corrected for right-censoring in the data when computing the fractions of spells that end within a given number of months into work and nonparticipation.

Finally, we can compare the estimated covariate effects of the structural model with the estimation results found in Section 4. The covariate effects in the wage offer distribution of the structural model largely coincide with those of the estimated random effects wage equation. The signs of the covariate effects always coincide and the sizes of the effects are close to each other. The covariate effects in the competing risks model are more difficult to compare. Except for the benefits level the individual characteristics have the same effect on the exit rate to nonparticipation in the reduced-form competing risks model as in the utility of nonparticipation in the structural model. Comparing the exit rate to work in the competing risks model with the job offer arrival rate in the structural model does not give many covariates with similar effects. Recall that the exit rate to work in the structural model does not only consist of the job offer arrival rate, but also includes the probability of accepting a job offer, which may cause the discrepancy in covariate effects between the exit rate to work in the reduced-form model and the job offer arrival rate in the structural model.

As mentioned earlier, the parameters of most interest are those describing duration dependence. Since duration dependence is only substantial and significant in the wage offer distribution our result provide evidence for only some theoretical explanations for the high degree of persistence in European unemployment. The negative duration in the wage offer distribution might imply that workers lose skills while being unemployed (e.g. Ljungqvist and Sargent, 1998). The absence of duration dependence in the job offer arrival rate makes stigmatization of long-term unemployed workers or discouragement less likely explanations for high levels of unemployment persistence (e.g. Blanchard and Diamond, 1994; Calmfors and Lang, 1995). The source of duration dependence is not the only relevant factor for targeting unemployment policies. Machin and Manning (1999) stress that also the timing is particularly important, active labor market policies should be aimed at the period where the negative duration dependence is most pronounced. Below we investigate how the duration dependence in the wage offer distribution affects reservation wages and the exit rate to work.

For some particular types of unemployed workers we show the estimated reservation wage path and probabilities of going to employment. We also show what these paths would be if we subsequently ignore (i) the option of entering nonparticipation, and (ii) the duration dependence in the wage offer distribution.[12] We define four types of unemployed workers. All of these types are 30 years, have the

[12] Since we do not find any substantia duration dependence in the job offer arrival rate, it is not particularly interesting to consider the case where duration dependence in the job offer arrival rate is ignored.

German nationality, entered unemployment after a period of employment, and receive 1000 DMark unemployment benefits. For the unobserved heterogeneity we assume that all four individuals are in the group with the most probability mass, i.e. they have both high job offer arrival rates and on average high wage offers. The first individual is a married man without children, who lives in South-Germany, and experienced 10 years of education. The second individual is a single mother living in South-Germany with 6 years of education. The third individual is again a married man without children, but with only 6 years of education. And the fourth individual is identical to the first individual, but he lives in East-Germany. Figures 4 to 11 show for each of these types of individuals the reservation wage paths and re-employment probabilities. It is clear that if the reservation wage is always above the utility of nonparticipation, then ignoring nonparticipation does not affect the reservation wage path and the exit rate to work. As can be seen from the figures this is the case for all types except the single mother.

The main result from the figures is that ignoring the duration dependence in the wage offer distribution biases both the estimated reservation wage and the re-employment probabilities. This stresses the importance of duration dependence in explaining the transition from unemployment to work. The reservation wage path crosses the instantaneous utility of nonparticipation only for the single mother. For the other three types of individuals the reservation wages are much higher than the utility of being nonparticipant. It is well known that, compared to males, a relatively large proportion of females leave unemployment to nonparticipation (e.g. Machin and Manning, 1999). This is particularly the case for single mothers, not only in Germany, but also in other OECD countries (see the OECD Employment Outlook, 1991, 1995, 1999). This also implies the relevance of taking nonparticipation into account when modeling the process of job search of this group of unemployed workers. Ignoring the option of entering nonparticipation would bias the estimated reservation wage path. The relevance of allowing for nonparticipation is less high for the other three specified types of individuals. Their behavior is only affected if the unobserved component in the utility of nonparticipation has an extreme value.

Our estimation results are largely in agreement with the assumptions of Ljungqvist and Sargent (1998). They argue that the persistence in European unemployment can be explained from a combination of the generous benefits system and loss of skills at the start of unemployment and while being unemployed, while the job offer arrival rate is assumed to be constant. Our results confirm these assumptions. Ljungqvist and Sargent (1998) define 21 skill levels on a linear scale and assume that while being unemployed each two weeks a worker has probability 0.2 that his skills drop 1 level. If workers who become unemployed are on average in the 10th skill level, this implies that after 3 months workers have lost on average 12% of their skills, after 6 months 24% of their skills, after one year 48% of their skills and after two years 82% of their skills. If we compare these percentages with our empirical findings we see that the expected wage associated to a job offer drops around 7% per three months during the first 15 months and around 4% per three months afterwards. After two years of unemployment workers have thus lost slightly less than 50% of their skills. One can think of a number of policies to stimulate re-employment. In line with Ljungqvist and Sargent (1998) lowering the unemployment benefits level could be an effective policy. However, it should be noted that the effect of the benefits level on re-employment has not been a focus of this paper. Alternatively, active labor market policies aimed at improving the prospects of unemployed workers should focus on improving the skills of the unemployed workers, or on avoiding a loss of skills amongst unemployed workers. Schooling and training programs or work experience jobs are therefore expected to be more successful than, for example, counseling or monitoring. Since the decline in the wage offer distribution is fastest in the first year of unemployment, active labor market policies should start early in the unemployment spell and thus not only focus on long-term unemployed workers.

The policy implications above depend on the loss of skills interpretation as the cause of the sharp negative duration dependence in job offers. There are several other possible reasons however for negative duration dependence during unemployment starting already in the first months after unemployment. First, individuals lose their bargaining position at the moment they become unemployed and their firm specific capital becomes worthless. An unemployed worker recalled by a former employer will probably not have lost the firm specific skills and can thus bargain a higher wage. Since recalls usually occur at the beginning of the unemployment spell, these bias the wage offer distribution upwards for short spells. Similarly, if at the start of the unemployment spell the individual knows that he will start a job within a short period, the job is the result of on-the-job search, which usually have higher bargained wages. However, such short periods of unemployment are not very likely to show up in our self-reported data, as these individuals will probably not classify themselves as being unemployed even though they receive unemployment benefits.

Another possibility is that at the beginning of unemployment individuals use their social network of employed friends and relatives to search for work. Since the information concerning job search in such social networks is limited, 'informal' job search is only efficient at the beginning of the unemployment spell. Indeed, Gregg and Petrongolo (1997) have shown that individuals lose contacts while being unemployed. This is also a loss of human capital, but not one for which it is clear how active labor market policies could compensate for.

6 Conclusions

In this paper we have argued that the transition from unemployment to nonparticipation is a choice instead of a stochastic occurrence. Unemployed workers are considered to permanently have the option of becoming nonparticipant, and to quit searching for work at the moment their reservation wage decreases below their utility of nonparticipation. Allowing for this transition improves the identification of duration dependence in the structural elements of the job search model.

Our estimation results show that during the unemployment spell the rate at which job offers arrive is relatively constant but that the wage offer distribution shifts downwards. The latter effect causes the reservation wage to decrease over the unemployment duration. We argue that this shift in the wage offer distribution is associated with loss of skills. Our empirical results thus confirm the assumption of workers losing skills during unemployment made by Ljungqvist and Sargent (1998). The policy implication is that active labor market policies aimed at improving the prospects of the unemployed workers should focus on improving skills. Schooling and training programs or subsidized employment are therefore expected to be more successful than, for example, counseling or monitoring. The programs should not only focus on the long-term unemployed workers as workers start losing skills already in the first year of unemployment. However, some human capital lost in the first few months of unemployment, such as social networks, might be hard to compensate for by schooling and training programs.

The most important factors in explaining the utility of nonparticipation are gender, having children and the region of residence. Females, unemployed workers with children and individuals living in former West-Germany have a higher utility of being nonparticipant. Additionally, 'years of education' is the most important covariate in both the job offer arrival rate and the wage offer distribution. Labor market prospects definitely improve with years of education.

Finally, our results indicate that both duration dependence and unobserved heterogeneity are relevant in explaining job search and the transition from unemployment to employment. This has been recognized in reduced-form studies of labor market transitions. However, most structural analyses of labor market behavior do not account for non-stationarity in job search. For particular groups of workers taking account of nonparticipation as an independent labor market state is important, as it changes the complete reservation wage path and the transition from unemployment to nonparticipation is endogenous and selective. This is particularly true for individuals who do not have very good labor market prospects and have a high utility of nonparticipation, for example single mothers.

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Appendix: The job search model

Our model consists of assumption 1-5 from Van den Berg (1990a), and adds the permanent option of going into nonparticipation. These assumptions are summarized below:

Assumption 1 *For each $t \geq 0$, the wage offer distribution has the following properties: (i) $F_t(w)$ is a continuous function, which is strictly monotonically increasing on the support $\langle \alpha_t, \beta_t \rangle$ (with $0 \leq \alpha_t < \beta_t \leq \infty$), (ii) $F_t(\alpha_t) = 0$, (iii) $\lim_{w \to \beta_t} F_t(w) = 1$, and (iv) the mean is bounded.*

Assumption 2 *For each $t \geq 0$, the job offer arrival rate λ_t satisfies $0 \leq \lambda_t \leq K$ for a fixed $K < \infty$. Also, the benefits level while being unemployed, b_t, satisfies $0 \leq b_t \leq K$.*

Assumption 3 *Except for only a finite number of points (possibly 0), $F_t(w)$, α_t, b_t and λ_t are continuous functions of t. If a function is discontinuous in a point, say at t^*, then it is right-continuous and the left-hand limit of this function at t^* exists.*

Assumption 4 *The functions $F_t(w)$, b_t and λ_t are constant on $[T, \infty\rangle$ for some positive T.*

Assumption 5 *The discount rate ρ satisfies $0 < \rho < \infty$.*

Assumption 6 *At each $t \geq 0$ an unemployed worker has the option of moving into the absorbing state of nonparticipation, where the instantaneous utility equals $u_{np,t}$. $u_{np,t}$ is continuous for all t and constant on $[T, \infty\rangle$ for some positive T.*

The optimal strategy of an unemployed worker is characterized by the following theorems:

Theorem 1 ***(Determination of the reservation wage path)*** *Under Assumption 1-6, the optimal strategy consists of a time to go into nonparticipation $\bar{t}$ and a reservation wage path $\phi_t(\bar{t})$ for $t \leq \bar{t}$, given $\bar{t}$. The reservation wage path follows the differential equation*

$$\frac{\partial \phi_t(\bar{t})}{\partial t} = \rho \phi_t(\bar{t}) - \rho b_t - \lambda_t \int_{\phi_t(\bar{t})}^{\infty} (w - \phi_t(\bar{t}))\, dF_t(w) \tag{4}$$

for each $t \leq \bar{t}$ where the functions λ_t, b_t, and $F(w;t)$ are continuous. If these functions are not continuous in t, then the right-hand side of (4) gives the right-hand derivative of ϕ_t in t. The left-hand derivative is calculated by replacing the functions in (4) by their left-hand limits at that discontinuity point. There holds $\phi_{\bar{t}}(\bar{t}) = \rho \int_{\bar{t}}^{\infty} u_{np,t} \exp(-\rho(t-\bar{t}))dt$ if $\bar{t} \leq T$ and $\phi_{\bar{t}}(\bar{t}) = r_T(r_T)$ if $\bar{t} = \infty$. Here $r_t(x)$ follows from $r_t(x) = b_t + \frac{\lambda_t}{\rho} \int_x^{\infty} (w-x)dF_t(w)$ for all t. Now, if $0 < \bar{t} < \infty$, then $\bar{t}$ should satisfy the condition

$$\lim_{t \uparrow \bar{t}} r_t(\phi_t(t)) \geq u_{np,\bar{t}} \geq \lim_{t \downarrow \bar{t}} r_t(\phi_t(t)) \tag{5}$$

Proof:
The reservation wage path in equation (4) directly follows from Van den Berg (1990a), where entering nonparticipation can be interpreted as accepting a job with certainty at $\bar{t}$ with wage $w = \rho \int_{\bar{t}}^{\infty} u_{np,t} \exp(-\rho(t-\bar{t}))dt$.

We prove condition (5) by noting that for points $\bar{t}$ where λ_t, b_t, and $F(w;t)$ are continuous, $\lim_{\Delta \downarrow 0} \frac{\phi_{\bar{t}-\Delta}(\bar{t}) - \phi_{\bar{t}-\Delta}(\bar{t}-\Delta)}{\Delta} = 0$. This implies that just before $\bar{t}$ the individual is indifferent between entering nonparticipation and searching for work until $\bar{t}$. The condition can be expanded because

$$\phi_{\bar{t}-\Delta}(\bar{t}) = \rho\Delta r_{\bar{t}-\Delta}(\phi_{\bar{t}}(\bar{t})) + (1-\rho\Delta)\phi_{\bar{t}}(\bar{t}) + \sigma(\Delta)$$

and

$$\begin{aligned}
\phi_{\bar{t}-\Delta}(\bar{t}-\Delta) &= \phi_{\bar{t}}(\bar{t}) - \Delta \frac{d\phi_{\bar{t}}(\bar{t})}{d\bar{t}} + \sigma(\Delta) \\
&= \phi_{\bar{t}}(\bar{t}) - \rho\Delta \frac{d \int_{\bar{t}}^{\infty} u_{np,t} \exp(-\rho(t-\bar{t}))dt}{d\bar{t}} + \sigma(\Delta) \\
&= \phi_{\bar{t}}(\bar{t}) - \Delta\rho(\phi_{\bar{t}}(\bar{t}) - u_{np,\bar{t}}) + \sigma(\Delta)
\end{aligned}$$

Solving for $r_{\bar{t}-\Delta}(\phi_{\bar{t}}(\bar{t}))$ and taking the limit of Δ to 0 gives $r_{\bar{t}}(\phi_{\bar{t}}(\bar{t})) = u_{np,\bar{t}}$.

For points where λ_t, b_t, and $F(w;t)$ are not continuous, an analogue reasoning implies we have to replace $r_{\bar{t}}(\phi_{\bar{t}}(\bar{t}))$ by the left and right-hand limits of $r_t(\phi_t(t))$ at $\bar{t}$, which gives condition (5). □

Condition (5) becomes simpler in the special case considered in this paper where $u_{np,t}$ is constant for all t. Since $\phi_{\bar{t}}(\bar{t}) = u_{np}$, the condition simplifies to $\lim_{t\uparrow\bar{t}} r_t(u_{np}) \geq u_{np} \geq \lim_{t\downarrow\bar{t}} r_t(u_{np})$, i.e. individuals go to nonparticipation when the instantaneous value of search equals the instantaneous utility of nonparticipation.

There might be no values or multiple values $\bar{t}$ that satisfy condition (5). To determine the optimal moment of entering nonparticipation we use the next theorem:

Theorem 2 ***(Selection of the optimum)*** *Let S denote the union of the set of all values $\bar{t}$ that satisfy condition (5) and $\{0, \infty\}$. The optimal time of entering nonparticipation satisfies*

$$\bar{t} = \arg\max_{s \in S} \phi_0(s)$$

Proof:
The proof of this theorem is straightforward. Out of a finite number of possible moments of entering nonparticipation, the unemployed worker chooses the moment which generates the highest present life-time utility. The points $\bar{t} = 0$ and $\bar{t} = \infty$ allow for the possibility that the unemployed worker enters nonparticipation immediately ($r_0(\phi_0(0)) < u_{np,0}$) and that the unemployed worker never enters nonparticipation ($r_T(r_T) > u_{np,T}$). $\square$

Under some conditions the set S includes additional to $\{0, \infty\}$ at most one possible $\bar{t}$. This is when $r_t(x)$ is non-increasing over time for any x and when $\int_t^{\infty} u_{np,s} \exp(-\rho(s))ds$ is non-decreasing for all t. Sufficient, but not necessary, conditions for this situation are:
- b_t and λ_t are non-increasing.
- $F_t(w)$ first-order stochastically dominates $F_{t+\tau}(w)$ for all t and $\tau > 0$.
- $u_{np,t}$ is non-decreasing.
In these circumstances there can be only one $\bar{t}$ that maximizes lifetime utility.

Exit Destination	Observed Work	Observed Nonp.	R-C	Total
Individual characteristics				
Age (in years)	30.2	29.3	30.7	
	(7.2)	(6.8)	(7.5)	
Male	73%	2%	25%	1306
Female	56%	13%	31%	1343
Children	63%	10%	27%	1803
No children	67%	3%	30%	846
Education (in years)	11.3	10.9	11.1	
	(2.4)	(2.3)	(2.3)	
Marital status missing	64%	9%	27%	277
Married	63%	10%	27%	1244
Single	67%	4%	29%	1128
German nationality	65%	7%	28%	2093
Non-German nationality	64%	9%	27%	556
Regions				
South-Germany	70%	8%	22%	566
Mid-Germany	68%	8%	23%	251
North-Germany	62%	11%	27%	651
East-Germany	62%	5%	33%	1031
Berlin	63%	7%	30%	150
Labor market state before inflow into unemployment				
Employment	67%	5%	28%	2132
Out of the labor force	55%	18%	27%	517
Benefits level				
(in D-Mark)	1077	888	1054	
	(564)	(495)	(563)	
Reservation wage				
Observed	54%	7%	39%	367
(in D-Mark)	1869	1554	1759	
	(700)	(539)	(555)	
Unemployment duration				
(in months)	6.0	9.0		
	(5.2)	(5.2)		
Post-unemployment wage				
Observed	100%			1237
(in D-Mark)	2499			
	(953)			
# spells	1710	200	739	2649
# individuals				1938

Explanatory note: We distinguish three types of spells: spells observed to exit into work, spells observed to exit into nonparticipation, and right-censored spells. The table shows how the spells of individuals with certain characteristics are distributed over the three types of spells. For the individual characteristics ‘Age’ and ‘Education’ and for ‘Benefits level’, ‘Reservation wage’, ‘Unemployment duration’ and ‘Post-unemployment wage’, we give the subsample means (and standard deviations). The last column provides the total number of spells of individuals in the sample with a certain characteristic.

Table 1: Some characteristics of the data set.

	Work θ_w		Nonparticipation θ_n	
Unobserved heterogeneity				
v^a	−1.94	(0.60)	−5.19	(1.18)
v^b	−2.85	(0.53)	−5.52	(1.21)
v^c	−4.37	(0.58)	−7.52	(1.58)
p_{aa}	0			
p_{ab}	0			
p_{ac}	0.14	(0.34)		
p_{ba}	0			
p_{bb}	0.58	(0.60)		
p_{bc}	0			
p_{ca}	0.28	(0.26)		
p_{cb}	0			
p_{cc}	0			
Individual characteristics				
log(Age/10)	−0.28	(0.17)	−1.08	(0.40)
Female	−0.73	(0.073)	1.91	(0.30)
Children	−0.074	(0.076)	0.62	(0.22)
log(Education)	0.89	(0.19)	−0.17	(0.43)
Married	0.12	(0.089)	0.75	(0.20)
Marital status unobserved	0.19	(0.16)	0.37	(0.35)
Non-German nationality	−0.34	(0.12)	0.043	(0.26)
Inflow after out of the labor force	−0.15	(0.086)	0.47	(0.16)
log(Benefits/100)	−0.14	(0.061)	−0.077	(0.16)
Regions				
South-Germany	0		0	
Mid-Germany	−0.21	(0.12)	−0.27	(0.30)
North-Germany	−0.37	(0.096)	−0.088	(0.22)
East-Germany	−0.73	(0.098)	−1.16	(0.23)
Berlin	−0.58	(0.16)	−0.62	(0.39)
Duration dependence (monthly)				
λ_{1--3}	0		0	
λ_{4--6}	−0.10	(0.070)	0.10	(0.25)
λ_{7--9}	−0.13	(0.093)	0.62	(0.24)
λ_{10--12}	−0.084	(0.11)	1.01	(0.24)
λ_{13--15}	−0.20	(0.14)	0.97	(0.27)
λ_{16--18}	−0.25	(0.16)	0.23	(0.38)
λ_{19--21}	−0.30	(0.19)	0.42	(0.41)
λ_{22--}	−0.37	(0.26)	−0.51	(0.77)
$\log \mathcal{L}$	−6799.33			
# individuals	1938			
# spells	2649			

Explanatory note: Standard errors in parentheses.

Table 2: Estimation results of the competing risks model.

Intercept		
α	6.81	(0.13)
Individual characteristics		
log(Age/10)	0.16	(0.055)
Female	−0.23	(0.021)
Children	−0.036	(0.023)
log(Education)	0.25	(0.052)
Married	−0.038	(0.027)
Marital status unobserved	0.036	(0.054)
Non-German nationality	0.016	(0.033)
Inflow after out of the labor force	−0.074	(0.024)
log(Benefits/100)	0.20	(0.019)
Previous unemployment duration (in months)	−0.0057	(0.0019)
Regions		
South-Germany	0	
Mid-Germany	0.026	(0.042)
North-Germany	−0.027	(0.030)
East-Germany	−0.27	(0.028)
Berlin	−0.0060	(0.046)
Standard deviations		
σ_v	0.18	(0.016)
σ_ε	0.26	(0.0089)
$\log \mathcal{L}$	809.30	
# individuals	1026	
# observations	1237	

Explanatory note: Standard errors in parentheses.

Table 3: Estimation results of the random effects wage equation.

	Job offer arrival rate λ		Wage offer distribution μ		Utility of nonparticipation u_{np}	
Duration dependence (month)						
0	0		0			
3	0.083	(0.23)	−0.076	(0.017)		
6	0.18	(0.27)	−0.14	(0.020)		
9	0.19	(0.32)	−0.21	(0.024)		
12	−0.070	(0.42)	−0.27	(0.028)		
15	0.054	(0.47)	−0.35	(0.033)		
18	−0.16	(0.55)	−0.37	(0.035)		
21	−0.028	(0.63)	−0.43	(0.043)		
24	0.085	(0.61)	−0.48	(0.051)		
Unobserved heterogeneity						
v_a	−11.6	(5.21)	5.33	(0.17)		
v_b	−3.53	(6.86)	5.75	(0.15)		
p_{aa}	0.093	(0.064)				
p_{ab}	0.18	(0.069)				
p_{ba}	0.018	(0.052)				
p_{bb}	0.71	(0.053)				
Individual characteristics						
Intercept					6.35	(0.56)
log(Benefits/100)					0.36	(0.089)
log(Age/10)	2.24	(1.75)	0.27	(0.051)	−0.15	(0.18)
Female	−3.60	(1.20)	−0.29	(0.024)	0.53	(0.26)
Children	1.51	(0.80)	−0.0073	(0.023)	0.32	(0.13)
log(Education)	3.98	(2.03)	0.44	(0.052)	0.091	(0.25)
Married	−0.90	(0.95)	−0.041	(0.025)	0.16	(0.11)
Marital status unobserved	3.62	(2.89)	0.020	(0.051)	0.12	(0.18)
Non-German nationality	0.37	(1.09)	0.026	(0.034)	0.061	(0.13)
Inflow after out of the labor force	−1.98	(0.75)	−0.041	(0.029)	0.34	(0.16)
Regions						
South-Germany	0		0		0	
Mid-Germany	1.41	(1.25)	0.0071	(0.042)	−0.074	(0.15)
North-Germany	0.70	(0.86)	−0.061	(0.030)	0.025	(0.12)
East-Germany	0.81	(0.83)	−0.34	(0.030)	−0.56	(0.16)
Berlin	−1.12	(1.30)	−0.044	(0.050)	−0.15	(0.17)
ρ (yearly)	0.19	(0.032)				
σ_w			0.48	(0.031)		
σ_{np}					0.54	(0.13)
Measurement errors						
η_w	0.26	(0.0067)				
η_ϕ	0.27	(0.011)				
$\log \mathcal{L}$	−19872					
# individuals	1938					
# spells	2649					

Explanatory note: Standard errors in parentheses.

Table 4: Estimation results of the structural model.

	Data		Reduced-form model		Structural model	
month	**work**	**nonpart.**	**work**	**nonpart.**	**work**	**nonpart.**
1	10.9%	0.6%	11.2%	0.5%	10.0%	0.5%
			(0.4)	(0.1)	(0.4)	(0.02)
2	20.8%	1.0%	20.3%	1.0%	18.7%	1.0%
			(0.7)	(0.2)	(0.6)	(1.0)
3	28.9%	1.4%	27.8%	1.5%	26.0%	1.5%
			(0.9)	(0.2)	(0.7)	(2.0)
4	34.6%	1.9%	33.5%	1.9%	32.4%	1.9%
			(0.9)	(0.3)	(0.8)	(1.4)
5	39.2%	2.1%	38.4%	2.3%	37.9%	2.4%
			(0.9)	(0.3)	(0.8)	(1.4)
6	44.2%	2.7%	42.7%	2.7%	42.6%	2.9%
			(1.0)	(0.3)	(0.8)	(1.7)
7	47.8%	3.2%	46.3%	3.3%	46.7%	3.4%
			(1.0)	(0.3)	(0.9)	(1.9)
8	50.9%	4.0%	49.5%	3.9%	50.3%	3.9%
			(1.0)	(0.4)	(0.9)	(3.2)
9	54.0%	4.5%	52.4%	4.4%	53.3%	4.3%
			(1.0)	(0.4)	(0.9)	(3.4)
10	56.3%	5.1%	55.1%	5.1%	56.0%	4.7%
			(1.0)	(0.4)	(0.9)	(2.4)
11	58.2%	5.5%	57.4%	5.8%	58.3%	5.1%
			(1.0)	(0.5)	(0.9)	(5.7)
12	61.2%	6.6%	59.5%	6.4%	60.3%	5.5%
			(1.1)	(0.5)	(0.9)	(3.0)
13	63.2%	7.3%	61.2%	6.9%	62.0%	5.8%
			(1.0)	(0.5)	(0.9)	(1.3)
14	64.5%	7.5%	62.7%	7.4%	63.4%	6.2%
			(1.0)	(0.5)	(0.9)	(1.4)
15	65.8%	8.0%	64.1%	7.9%	64.6%	6.5%
			(1.1)	(0.6)	(0.9)	(2.1)
16	67.0%	8.1%	65.3%	8.1%	65.7%	6.6%
			(1.0)	(0.6)	(0.9)	(1.9)
17	68.2%	8.3%	66.4%	8.3%	66.7%	6.7%
			(1.0)	(0.6)	(0.9)	(4.5)
18	69.3%	8.5%	67.5%	8.4%	67.6%	6.9%
			(1.1)	(0.6)	(0.9)	(1.3)
19	70.2%	8.8%	68.4%	8.6%	68.4%	7.0%
			(1.0)	(0.6)	(0.9)	(1.7)
20	71.1%	8.9%	69.2%	8.8%	69.0%	7.2%
			(1.0)	(0.6)	(0.9)	(0.8)
21	71.8%	9.1%	70.0%	9.0%	69.7%	7.3%
			(1.0)	(0.6)	(0.9)	(2.7)
22	72.3%	9.1%	70.7%	9.1%	70.2%	7.4%
			(1.0)	(0.6)	(0.9)	(3.0)
23	73.2%	9.2%	71.4%	9.2%	70.7%	7.5%
			(1.0)	(0.6)	(0.9)	(4.1)
24	73.9%	9.2%	72.0%	9.2%	71.1%	7.6%
			(1.1)	(0.6)	(0.9)	(1.5)

Explanatory note: standard errors in parentheses.

Table 5: Percentage of spells that end in work and nonparticipation within a given number on months in the data, according to the reduced-form competing risks model and according to the structural model.

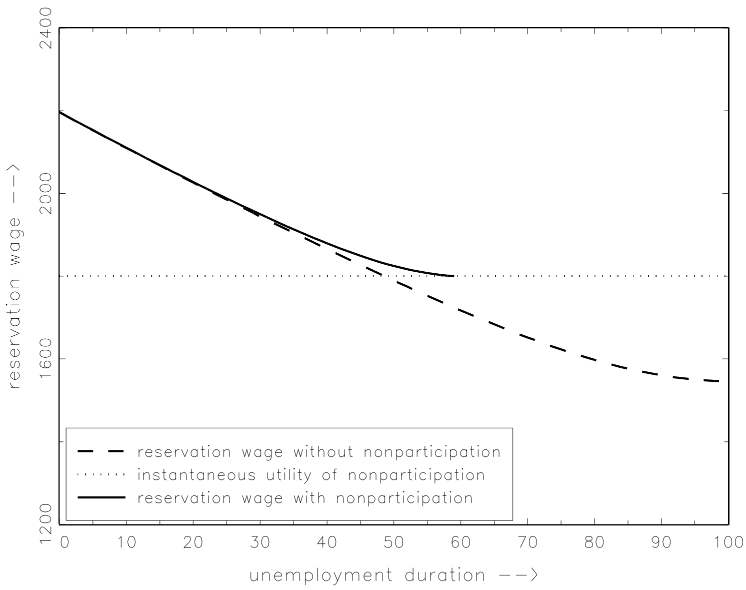

Figure 1: The reservation wage path with and without allowing for a permanent option of entering nonparticipation.

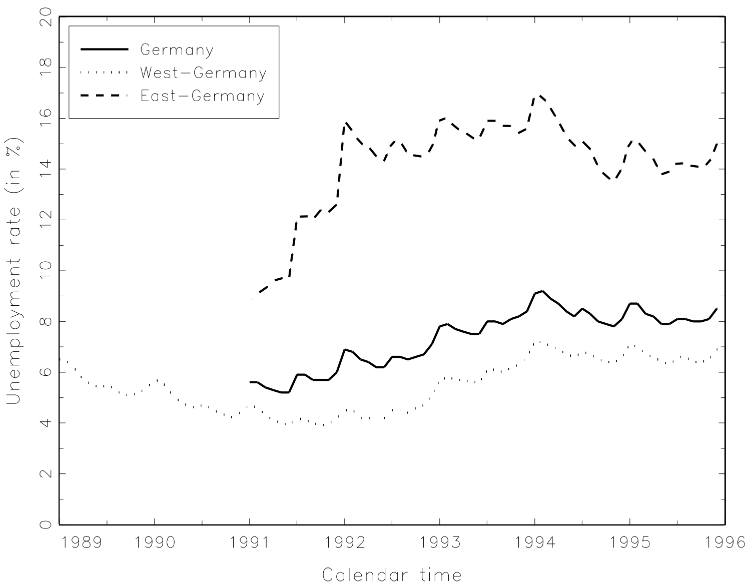

Figure 2: The monthly unemployment rate in Germany and stratified to former West-Germany and East-Germany. For East-Germany reliable unemployment figures only came available in 1991. (Source: Eurostat)

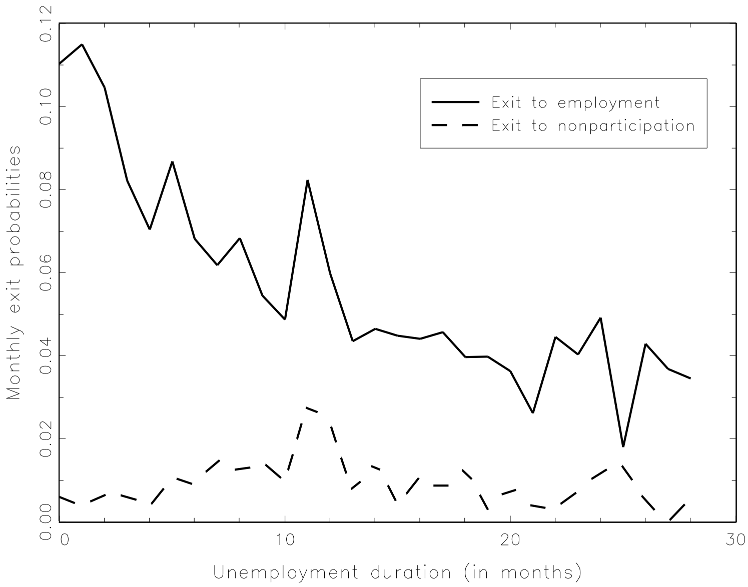

Figure 3: The empirical hazard rates from unemployment to employment and nonparticipation.

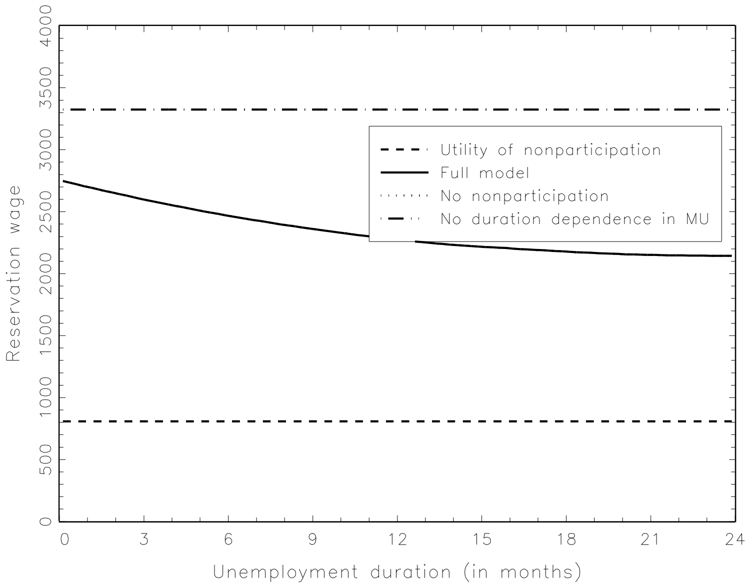

Figure 4: The estimated reservation wage path (man, 30 years, married, no children, 1000 DMark benefits, 10 years of education, living in South-Germany, inflow from employment, German nationality).

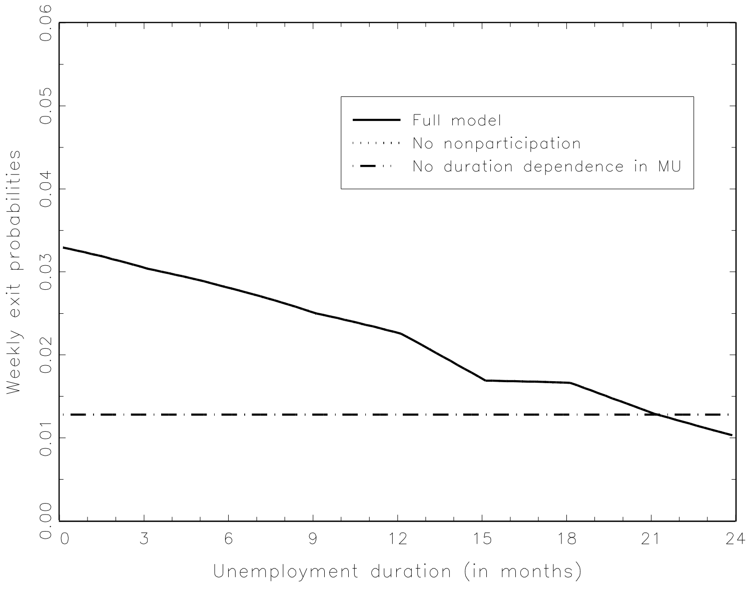

Figure 5: The estimated re-employment probability (man, 30 years, married, no children, 1000 DMark benefits, 10 years of education, living in South-Germany, inflow from employment, German nationality).

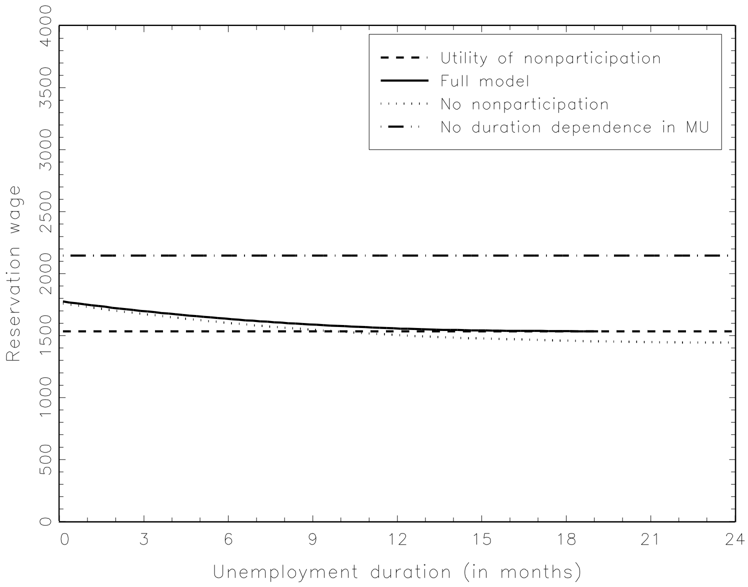

Figure 6: The estimated reservation wage path (woman, 30 years, single, children, 1000 DMark benefits, 6 years of education, living in South-Germany, inflow from employment, German nationality).

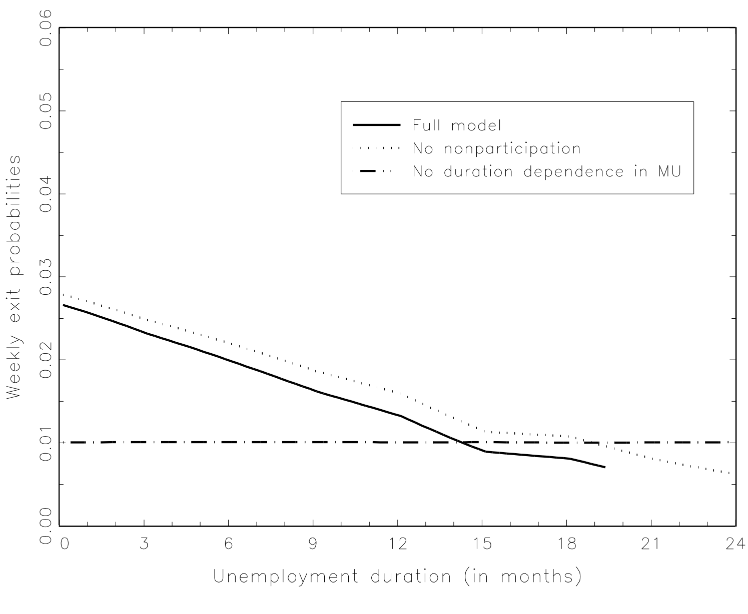

Figure 7: The estimated re-employment probability (woman, 30 years, single, children, 1000 DMark benefits, 6 years of education, living in South-Germany, inflow from employment, German nationality).

Figure 8: The estimated reservation wage path (man, 30 years, married, no children, 1000 DMark benefits, 6 years of education, living in South-Germany, inflow from employment, German nationality).

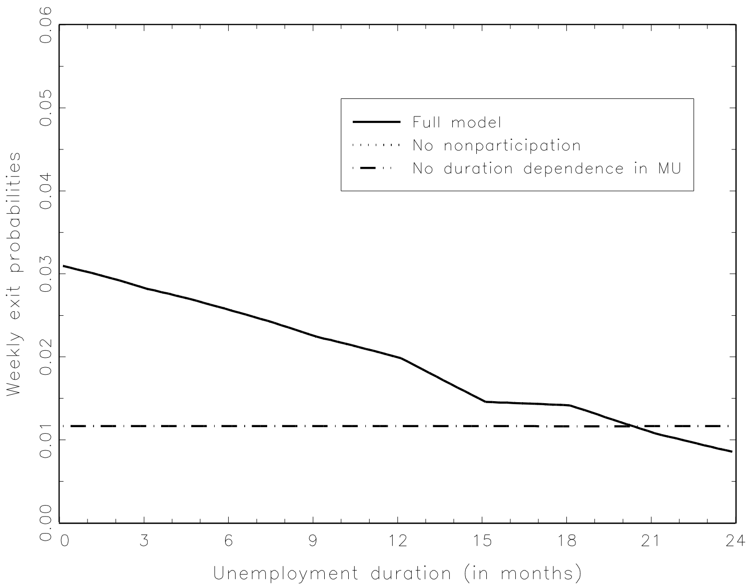

Figure 9: The estimated re-employment probability (man, 30 years, married, no children, 1000 DMark benefits, 6 years of education, living in South-Germany, inflow from employment, German nationality).

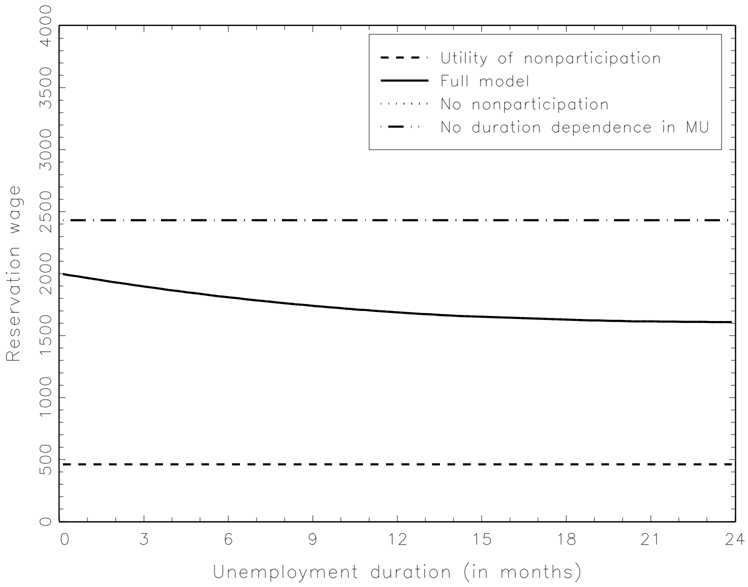

Figure 10: The estimated reservation wage path (man, 30 years, married, no children, 1000 DMark benefits, 10 years of education, living in East-Germany, inflow from employment, German nationality).

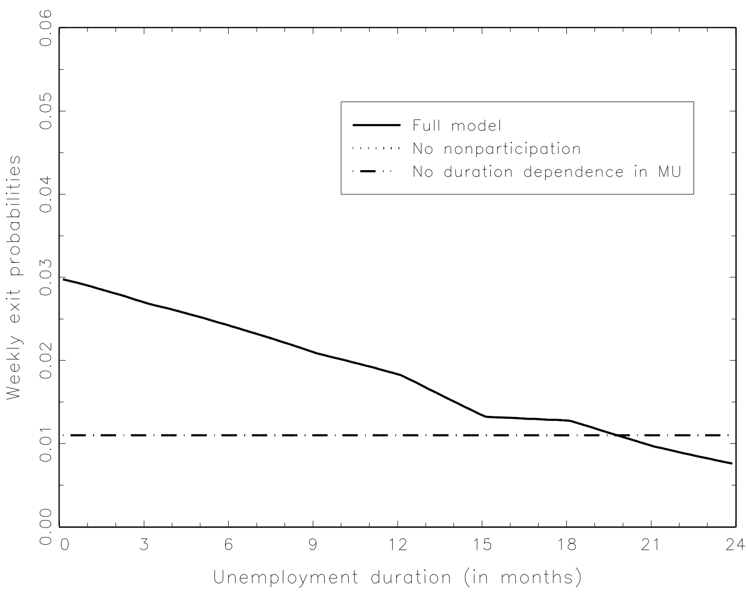

Figure 11: The estimated re-employment probability (man, 30 years, married, no children, 1000 DMark benefits, 10 years of education, living in East-Germany, inflow from employment, German nationality).